Applied Economics, 2006, **38**, 573–586

Risk and farm operator labour supply

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This study uses a large increase in US Federal crop insurance subsidies as a natural experiment to identify the importance of risk for farm operator labour supply. Subsidy increases induced greater crop insurance coverage, which in turn reduced farmers' financial risks. Crop insurance participation data are merged with farm-level Census of Agriculture data from 1992 and 1997 to compare how individuals' off-farm labour supply changed in response to the policy-induced change in insurance coverage. The empirical approach controls for unobserved heterogeneity and accounts for the censored nature of the data. It is found that greater insurance coverage reduces the off-farm labour supply of operators who produced at least $100 000 of output, and increased the labour supply of small-farm operators who produced less than $25 000 of output.

I. Introduction

Understanding how farm households respond to risk is important in the context of large government expenditures on programmes designed to mitigate farm business income risk. In the USA in 2003, more than one hundred crops and two-thirds of all cropland were covered by a federal crop insurance contract, with total indemnity payments of $3.24 billion. Crop insurance is one of the largest US government insurance programmes in terms of payments to individuals, with disbursements averaging $2620 per insured farmer in 2003, a large sum compared to $350 per covered worker for the Workers' Compensation programme in 1999 (RMA, 2004; Mont *et al.*, 2001). In addition to crop insurance, agricultural programmes are often designed, in part, to reduce farm income risk – through price supports, direct income support, or disaster relief. In 1999, 2000, 2001 and 2003, US government payments to farmers exceeded $20 billion, with average payments exceeding $15 000 per farm household.

Despite large expenditures on agricultural programmes to reduce farm income risk, it is unclear whether these programmes increase economic efficiency. Evidence that farmers alter their production decisions in response to risk would suggest that risk-coping is imperfect and potentially inefficient. This study uses a large increase in US Federal crop insurance subsidies as a natural experiment to identify the importance of business income risk for farm operators' labour supply. Subsidy increases induced greater crop insurance coverage, which reduced farmers' financial risks. We estimate operators' off-farm labour supply response to the policy-induced change in insurance coverage. Results of the analysis should indicate whether farmers made potentially inefficient production decisions to cope with risk despite the availability of private risk coping mechanisms, crop insurance, and large government expenditures on agricultural programmes.

This research is subsumed in a body of work examining how income risk associated with wages, unemployment, interest rates and health affects

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Applied Economics ISSN 0003–6846 print/ISSN 1466–4283 online © 2006 Taylor & Francis
http://www.tandf.co.uk/journals
DOI: 10.1080/00036840500369043

labour supply (Bell *et al.*, 1997; Basu *et al.*, 2001; Bingley and Walker, 2001; Berloffa and Simmons, 2003; Pistaferri, 2003; Coile, 2004). These studies attempted to measure how individuals or their spouses respond to these risks by adjusting hours worked, entering or exiting the work force, or finding a second job. Other related research has examined the efficacy and unintended consequences of public policies designed to mitigate labour income risk, focusing on how unemployment insurance or disability insurance alters labour market decisions (Capen *et al.*, 1985; Kaestner and Carroll, 1997; Cullen and Gruber, 2000; Gruber, 2000; Jurajda, 2002; Meyer, 2002).

While a few studies have examined whether workers use second or multiple jobs as hedges against unemployment or as ex post responses to negative financial shocks (Bell *et al.*, 1997; Boheim and Taylor, 2004), this study considers how small business income risk affects a sole proprietor's decision to work in a second job. Compared with workers, small business owners are likely to have more variable income and face the additional financial risk of business failure. Usitalo (2001) found that attitudes towards risk were important in the decision to become self-employed – with less risk-averse individuals more likely to become entrepreneurs. In this study we ask whether US farm business operators' respond to changes in their business risk by adjusting their off-farm labour supply. The agricultural sector is well suited for studying business owners' response to risk because farms produce a relatively homogeneous product and face a large amount of risk. In addition, off-farm employment is important for many US farm households (Sumner, 1982; Hallberg *et al.*, 1991; Mishra *et al.*, 2002). For the last three decades, about half of all farm operators worked some time off the farm, and about one third worked at least 200 days per year off-farm.

Empirical studies that have examined the relationship between risk and off-farm labour have used farm income variability as a measure of risk. Using a cross-sectional sample of Kansas farms, Mishra and Goodwin (1997) found that the coefficient of variation of gross farm income is a significant determinant of the operator's (but not the spouse's) supply of off-farm labour. Using county-level panel data from two states, Mishra and Goodwin (1998) found the coefficient of variation of farm income had a statistically significant and positive effect on off-farm labour supply. Kanwar (1999) used panel data on farms in India to estimate a two-stage labour supply model. Kanwar found the standard deviation of net returns has a significant positive effect on the decision to work, but does not significantly alter labour supply conditional on labour force participation. In another study of Indian farmers, Maitra (2001) found that small-scale farmers substituted off-farm labour for on-farm labour in response to idiosyncratic income shocks, while medium and large-scale farmers were able smooth consumption using credit markets.

This study takes a new approach to identify the importance of risk for off-farm labour supply by testing whether farm operators altered their labour allocation decisions in response to a large increase in crop insurance subsidies. In 1994, the US Congress passed the Federal Crop Insurance Reform Act (FCIRA), which markedly increased subsidies for premiums paid for crop insurance – fully subsidizing low levels of insurance (catastrophic coverage) and partially subsidizing higher levels of insurance. The increase in subsidies induced greater participation in crop insurance programmes and reduced some farmers' financial risks. The exogenous subsidy change allows us to compare changes in off-farm labour supply for farms facing different changes in insurance coverage. This 'difference in differences' comparison allows us to observe changes in off-farm labour supply in response to changes in risk while holding factors associated with the farm household constant.

Our approach differs from most previous work in two important respects. First, an exogenous instrument is used for risk – a large policy-induced decline in the cost of crop insurance. Most studies have used estimates of income variation to measure risk. A problem with this approach is that income variation is, to some extent, endogenous: farmers can adjust income variation by altering crop mix, fertilizer and pesticides applications, by purchasing an irrigation system, or altering their labour supply. If income variance is endogenous then the estimated relationship between income variance and labour allocation may be biased.

Second, unobserved individual heterogeneity is controlled for by examining differences in labour supply of the same farmers across time. Earlier studies used cross-sectional data or panel data with no individual fixed effects. These approaches cannot control for unobserved factors that might be correlated with income risk and labour allocation decisions. Such unobserved factors might include idiosyncratic agronomic variables such as soil types, cropping patterns and climate; prices of inputs and outputs; or characteristics and preferences of the farm operators. Methods that do not control for individual heterogeneity may falsely attribute differences in off-farm labour supply to differences in the riskiness of farm income rather than differences in agronomic factors associated with income risk. For example, consider two observationally identical farms, where cropping patterns are unobservable to the econometrician.[1] Assume one farm grows a crop with little net revenue variability that requires a lot of labour to produce, while a second farm grows a crop with greater net revenue variability that requires less labour to produce. Even if the farmers are risk neutral, the farmer of the risky crop should supply more labour off-farm, simply because the crop is less labour-intensive. In this case, the observed correlation between on-farm income risk and off-farm labour supply is spurious.

The next section provides a theoretical framework for considering how the off-farm labour supply of a farm operator may be influenced by a policy that alters the variance of uninsured on-farm income. Section III provides a brief background to the Federal Crop Insurance Reform Act and presents some descriptive statistics that illustrate the large changes in insurance coverage that resulted from the Act. Section IV discusses the empirical approach used to examine the effect of insurance rate subsidies on labour allocation decisions. Section V describes the data and measures of crop insurance coverage used in the study. The results of the empirical approach are discussed in Section VI and Section VII concludes the paper.

II. Theoretical Framework

Economic theory maintains that risk neutral individuals allocate their endowment of labour to equate the value marginal product of labour on-farm with the wage rate off-farm (e.g., Kerachsky, 1977; Sumner, 1982). If farmers are risk-averse and unable to smooth their income using capital or insurance markets, they might reduce the variation in their total income by reallocating labour from risky employment on-farm to less risky employment off-farm. While risk-averse farmers might work more off-farm to reduce their risk, in general, the direction of off-farm labour response to risk is ambiguous. Whether greater risk results in more or less off-farm labour depends on the manner in which risk enters the production function (Fabella, 1989; Kanwar, 1999; Hartwick, 2000) and how a farmer's marginal utility of income changes with the amount of leisure consumed. This section uses a simple model to show that a crop insurance policy change similar to FCIRA has an ambiguous effect on off-farm labour supply.

Consider a farm household model wherein the farm operator chooses an allocation of labour to maximize expected utility. The operator's labour endowment $\bar{L}$ is allocated to the farm f, off-farm employment l, or to leisure L, so $L=\bar{L}-f-l$. Let the operator's utility be a function of income and leisure, $U=U(I,L)$, where income I is derived from farm profits $\pi(f,p,\varepsilon)$, which is random, and off-farm income wl, which is not random. The variance of farm profits depends on an exogenous policy variable p (such as subsidized crop insurance) and an exogenous random factor ε. Define expected farm profits and the variance of farm profits as:

$$\mu = \mu(f,p) = E[\pi(f,p,\varepsilon)|p,f]$$

and

$$\sigma^2 = \sigma^2(f,p) = E\left[(\pi(f,p,\varepsilon) - m(f))^2 | p,f\right]$$

respectively. Let U^* be a second-order Taylor approximation of utility around the mean value of profits:

$$U \cong U^* = U + U_I(\pi - \mu) + \frac{1}{2} U_{II}(\pi - \mu)^2$$

where U, U_I, and U_{II} are evaluated at $\mu + wl$.[2] Throughout the presentation, subscripts indicate first or second order partial derivatives (e.g., $U_I = \partial U/\partial I$ and $U_{II} = \partial^2 U/\partial I^2$). Farmers are assumed to maximize $E[U^*]$, which simplifies to:

$$E[U^*] = U + \frac{1}{2} U_{II}\sigma^2 \qquad (1)$$

Maximizing Equation 1 with respect to f and l, and dropping third order terms produces the first order conditions:[3]

$$U_I\mu_f + \frac{1}{2} U_{II}\sigma_f^2 - U_L = 0$$
$$U_I w - U_L = 0$$

It can be seen from the first order conditions that if risk is additive, $\pi(f,p,\varepsilon) = \mu(f,p) + \varepsilon$, then $\sigma_f^2 = 0$,

[1] Although cropping patterns are theoretically observable, they are not always available in data and have not been used in some past studies. Even if observable, cropping patterns are also endogenous and tend to be closely tied to climate, soil type, water availability, and other exogenous factors associated with location. In earlier studies, risk measures have also been tied to location, so risk and other factors influencing production decisions may be confounded.

[2] The chain rule is used to take a Taylor expansion only around profits, because profits are the only random component of utility.

[3] Third order terms are dropped because they are likely to be small and because higher-order elements in the Taylor approximation were ignored.

which implies labour allocation decisions are separable from profit risk (as shown by Fabella, 1989). Risk could matter for labour decisions if σ_f^2 is non-zero. An assumption consistent with positive σ_f^2 is multiplicative farm income risk: $\pi(f, p, \varepsilon) = \varepsilon\mu(f, p)$. This assumption makes sense with a production function with limited substitutability, so that all factors scale up with labour. However, if production factors besides labour are constrained then on-farm labour may reduce risk on the margin and the sign of σ_f^2 would be negative.

Totally differentiating the first order conditions with respect to the policy variable p and solving the resulting system of equations for df/dp and dl/dp gives:

$$df/dp = \frac{-\left(\frac{1}{2}U_{II}\sigma_{fp}^2 + U_I\mu_{fp}\right)\left(U_{II}w^2 - U_{LI}w\right)}{\det H} \quad (2)$$

$$dl/dp = \frac{\left(\frac{1}{2}U_{II}\sigma_{fp}^2 + U_I\mu_{fp}\right)\left(U_{II}\mu_f w - U_{LI}w\right)}{\det H} \quad (3)$$

The denominators of Equations 2 and 3 (equal to the determinant of the Hessian of U^*) are assumed positive, a sufficient condition for a maximum. If the conditions $U_I > 0$, $U_L > 0$, $U_{II} < 0$, $U_{LL} < 0$ and $U_{LI} \geq 0$ hold then the term in the second parentheses in both Equations 2 and 3 is negative and the effect of the policy on labour allocations depends only on the sign of the term in the first parentheses. These assumptions are standard, except for U_{LI}. It seems reasonable that the marginal utility of leisure increases with income (a vacation is more enjoyable if one has more money to spend on it). If U_{LI} is sufficiently negative, however, the signs of Equations 2 and 3 may reverse. The remainder of this section assumes the terms in the second parentheses of both equations are negative.

The sign of the first parentheses in Equations 2 and 3 depends on sign and magnitude of $U_{II}\sigma_{fp}^2$ and $U_I\mu_{fp}$. A common assumption is that a policy change that decreases risk (such as subsidized crop insurance) reduces the marginal effect of labour on risk (i.e., if $\sigma_p^2 < 0$ then $\sigma_{fp}^2 < 0$). It is not obvious, however, that this is necessarily true. The term $U_I\mu_{fp}$, which measures the policy's effect on the average marginal utility of labour, is less ambiguous: if $\mu_p > 0$ (profits are subsidized) then it seems reasonable that $\mu_{fp} > 0$. For a policy that reduces farm income variance but does not affect average marginal profitability of labour ($\mu_{fp} = 0$), Equation 2 will have the same sign and Equation 3 will have the opposite sign as σ_{fp}^2. In other words, if $\sigma_{fp}^2 < 0$ then the policy results in more labour allocated on-farm and less off-farm.

In sum, it has been shown that a policy that reduces farm income variance has an ambiguous effect on labour allocation. The empirical section of this study estimates the relationship between insurance coverage and off-farm labour supply. Changes in insurance coverage between 1992 and 1997 depended in large part on the exogenous policy change in insurance premium subsidies. This theoretical section has shown that, in general, it is not possible to predict how an exogenous change in insurance coverage will affect off-farm labour supply.

III. Background

Beginning in the 1995 crop season, the Federal Crop Insurance Reform Act of 1994 (FCIRA) modified the federal crop insurance programme by authorizing the USDA to offer essentially 'free' catastrophic coverage to producers who grow an insurable crop.[4] Catastrophic coverage insures production losses falling below 50% of expected yield, indemnified at 55% of the expected market price of the insured crop. The FCIRA allows farmers to purchase additional coverage that provides a higher yield or revenue protection, with the premium on this 'buy-up' coverage subsidized by the government. For buy-up coverage, producers pay only a portion of the actuarial premium plus a small administrative fee. The share of the total premium paid by the government varies by coverage level. In 1997, the typical premium subsidy share was 42% on the 65% buy-up coverage.

The FCIRA had a large effect on the number of acres insured and the level of coverage (as measured by premiums). Figure 1 shows total subsidies, total premiums, and total acres enrolled in the crop insurance programme from 1990 to 1998. The total premium equals total farmer contributions plus total government subsidies. The figure presents separate plots for all crops and for the three largest individual crops (in acreage): corn, soybeans, and wheat. In 1997, these three crops accounted for 78.9% of the acreage insured, 55.5% of the subsidies, 51.7% of the total premiums, and 53.8% of cultivated cropland (excluding hay). The figure illustrates the marked increase in crop insurance coverage following implementation of the FCIRA and suggests

[4] The premium on this level of coverage is fully subsidized by the government but farmers must pay a $50 per crop per county administrative fee.

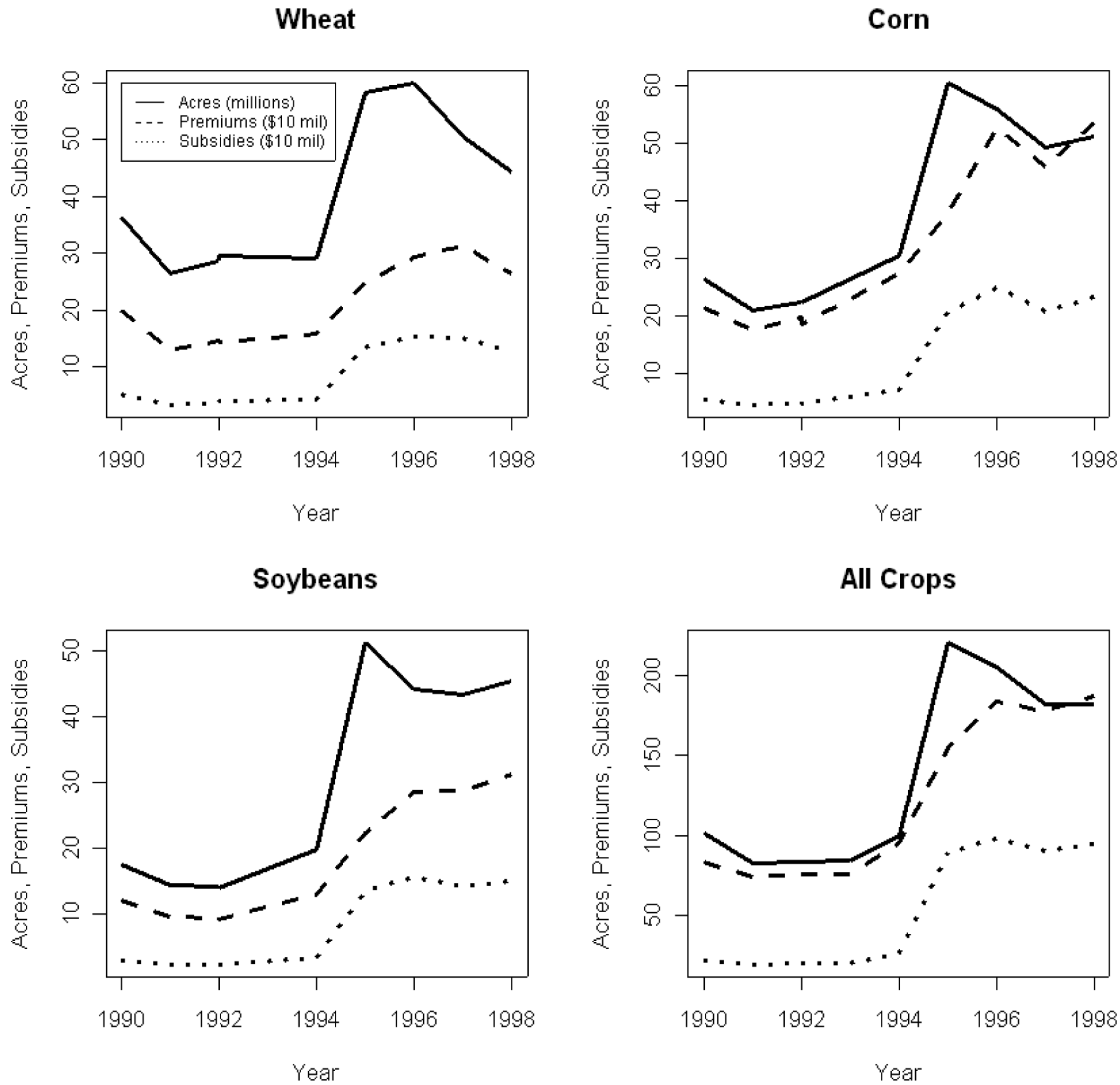

Fig. 1. Insurance coverage of all crops and largest individual crops in years preceding and following the FCIRA of 1994
Source: Risk Management Agency, at http://www.rma.usda.gov/data/

that the bulk of this increase stemmed directly from the increase in subsidies.

Table 1 gives additional information on the FCIRA for the ten crops that accounted for 85% of premiums paid in 1997. The table reports 1992 and 1997 levels of premiums, acres harvested, share of acres insured, premiums per acre harvested, premiums per insured acre, and subsidies per insured acre. There were large increases in premiums for most crops between 1992 and 1997. For barley, potatoes, and dry beans, premiums per acre harvested increased by about one-third; for wheat and sorghum, premiums increased by about one-half; and cotton, corn, and soybean premiums increased by almost two-thirds. The most extreme cases were peanuts, which showed little increase (the crop was heavily insured before the policy change), and tobacco, for which no federal crop insurance was available in 1992.

IV. Methods

This section outlines the approach used to identify the effect of crop insurance coverage on operator labour allocation. We begin by presenting an empirical model that relates off-farm labour supply to operator and farm characteristics and accounts for the censored nature of the data.

For farm operator i in time t, let the desired off-farm labour supply L_{it}^* be a function of factors X_{it} that influence the propensity to supply labour off farm, including wages, prices and characteristics

Table 1. Insurance coverage before and after FCIRA of 1994

	Total premiums ($1000)		Total acres harvested (1000)		Share of acres insured		Average premium per acre harvested ($/acre)		Average subsidy per acre insured ($/acre)		Average premium per acre insured ($/acre)	
	1992	1997	1992	1997	1992	1997	1992	1997	1992	1997	1992	1997
Wheat	146 118	313 933	59 003	60 953	0.497	0.833	2.53	5.16	1.36	2.98	5.09	6.2
Cotton	90 657	252 676	11 742	13 787	0.371	0.835	7.86	18.36	6.22	12.84	21.21	21.98
Corn	196 412	460 662	68 905	70 371	0.327	0.702	2.87	6.55	2.23	4.18	8.78	9.34
Dry beans	13 326	25 136	1159	1530	0.628	0.848	11.57	16.47	5.15	9.56	18.43	19.42
Sorghum	24 974	44 788	10 336	8351	0.351	0.755	2.45	5.38	1.96	3.59	6.98	7.13
Peanuts	39 840	36 153	1354	1292	0.780	0.914	29.54	28.01	8.77	13.67	37.86	30.63
Soybeans	93 715	288 374	54 672	66 135	0.262	0.659	1.74	4.37	1.69	3.29	6.62	6.63
Potatoes	12 497	28 857	905	1107	0.326	0.626	15.91	26.52	11.68	23.55	48.73	42.35
Barley	17 486	23 708	6463	5893	0.474	0.763	2.78	4.06	1.55	2.61	5.86	5.32
Tobacco	0	31 768	783	806	0	0.826	0	68.66	0	31.17	0	83.15

Source: Risk Management Agency at http://www.rma.usda.gov/data/

of the utility and production functions (Sumner, 1982; Huffman and Lange, 1989; Lass and Gempesaw, 1992; Benjamin *et al.*, 1996; Goodwin and Holt, 2002).[5] The propensity to supply off-farm labour may depend on the operator's per-acre crop insurance coverage C_{it} (defined in the next section), and this effect may depend on the scale of the operation:

$$L_{it}^* = \alpha + \beta X_{it} + \gamma C_{it} \cdot S_{it} + \beta_S S_{it} + \varepsilon_{it} \qquad (4)$$

where S_{it} is a dummy variable indicating the farm's scale category, α, β, γ, and β_S are parameters to be estimated, and ε_{it} is a random error. Per-acre coverage is used rather than total coverage in Equation 4 because total coverage is simultaneously determined with labour supply – total coverage (premium per acre times total acres) depends on farm size and farm size is closely related on-farm labour demand. Hence, one expects a negative relationship between total coverage and off-farm labour supply regardless of whether risk influences labour supply.

As discussed in the next section, farm-level data for this study are from the microfiles of the 1992 and 1997 US Census of Agriculture. In the census, farm operators categorically report the number of days they worked off-farm as 0, 1–49, 50–99, 100–49, 150–99 or 200 or more. Figure 2 illustrates the distribution of off-farm labour in 1992 and 1997 for the sample of 107 529 farms. As shown in the figure, farm operator off-farm labour is concentrated at two points: no off-farm labour days and 200-or-more days off farm. The concentration at these two points suggests that the 'desired' number of days supplied off-farm is censored below zero and above 200.

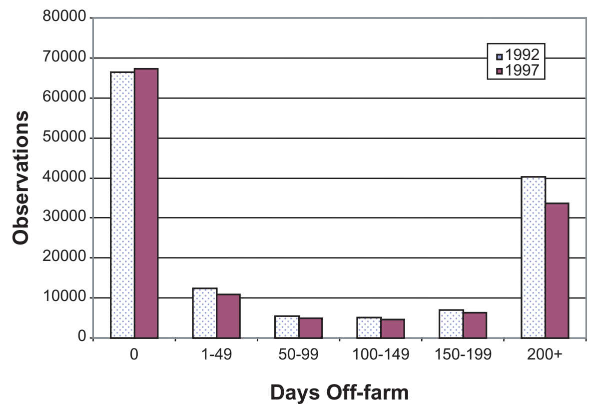

Fig. 2. Days off-farm in 1992 and 1997
Source: Census of Agriculture, 1992 and 1997. Observations = 107 529.

In other words, if the desired number of days supplied by operator i in time t (L_{it}^*) is less than zero, then the observed labour supply (L_{it}) is zero; and if the desired number of days supplied off-farm is greater than or equal to 200, one observes 200. Desired labour is not censored if it falls between 0 and 200 days:

$$L_{it} = \begin{cases} 0 & \text{if } L_{it}^* \leq 0 \\ L_{it}^* & \text{if } 0 < L_{it}^* < 200 \\ 200 & \text{if } 200 \leq L_{it}^* \end{cases} \qquad (5)$$

Some past labour market studies that have used county-level census data have treated the quantity of labour supplied off-farm as a continuous variable and have used the midpoints of the intervals as the

[5] The farm operator's off-farm labour supply may be a joint decision with the operator's spouse. However, there is no information about the operator's spouse so it is not included in the analysis.

observed quantity. A potential problem with this approach is that average labour supplied, conditional on being in a particular labour category, may not equal the midpoint of the category, resulting in biased estimates. To address this problem information is used from a representative national survey (ARMS) conducted in the census years to estimate the conditional means of each off-farm labour category.[6]

operator, the desired change in labour supply falls into one of three mutually exclusive categories: (1) *uncensored* – the operator's desired supply of labour is observed in both periods; (2) *unobservable* – the operator supplies no labour in both periods or supplies the maximum possible in both periods; or (3) *left or right censored* at the level of 'observed' change in the supply, $L_{i1} - L_{i0}$. That is, one observes:

$$\Delta L_i = \begin{cases} \Delta L_i^* \text{ (uncensored)} & \text{if } 0 < L_{i0}^* < 200 \text{ and } 0 < L_{i1}^* < 200 \\ \text{unobservable} & \text{if } (L_{i0}^* \leq 0 \text{ and } L_{i1}^* \leq 0) \text{ or } (L_{i0}^* \geq 200 \text{ and } L_{i1}^* \geq 200) \\ L_{i1} - L_{i0} \text{ (right censored)} & \text{if } (L_{i0}^* \leq 0 \text{ and } L_{i1}^* > 0) \text{ or } (L_{i0}^* < 200 \text{ and } L_{i1}^* \geq 200) \\ L_{i1} - L_{i0} \text{ (left censored)} & \text{if } (L_{i0}^* > 0 \text{ and } L_{i1}^* \leq 0) \text{ or } (L_{i0}^* \geq 200 \text{ and } L_{i1}^* < 200). \end{cases} \quad (7)$$

Identification based on differences

A problem with estimating the model defined by Equations 4 and 5 is that per-acre coverage is endogenous – farms make labour allocation decisions and insurance coverage decisions simultaneously, so the relationship between the two may not be causal. In this case the cross-sectional variance in pre-FCIRA coverage is most problematic. To identify the effect of insurance coverage on labour allocation, an attempt is made to remove variation in coverage not caused by the policy change. Because coverage was relatively low and grew by so much in response to the policy change, one can be confident that the policy change was the driving force for the *growth* in insurance coverage. The study, therefore, estimates how the labour supply of individual farms changed between periods in response to changes in the exogenous variables and changes in insurance coverage. Subtracting the first time period ($t = 0$) from the second ($t = 1$), one has:

$$\Delta L_i^* = \tilde{\alpha} + \tilde{\beta}\Delta X_i + \tilde{\gamma}\Delta C_i \cdot S_{i0} + \tilde{\beta}_S S_{i0} + \varepsilon_{it} \quad (6)$$

where S_{i0} is a dummy variable for the average farm scale category in the first period.

The censoring for the difference equation is complicated, as the dependent variable is the difference of a double-censored variable. The desired change in labour supplied off-farm ΔL_{it}^* is censored depending on the censoring of the desired off-farm labour in time zero L_{i0}^* and in time one L_{i1}^*. For each

In addition to controlling for aggregate changes that affected the labour supply of farms, estimating the difference regression (Equation 6) also controls for omitted variable bias. To illustrate, suppose the error term in Equation 4 contains time invariant factors that are correlated with the regressors: $\varepsilon_{it} = u_i + v_{it}$. These unobservable factors may include prices of inputs and outputs; specific features and locations of the land on which the farms are situated, such as soil types and climate; and characteristics and preferences of the farm operators. If the omitted variables are correlated with the regressors then estimates of Equation 4 will be biased. For example, the labour intensity of the crops grown may be positively correlated with coverage levels, which results in an inverse correlation between coverage and off-farm labour, causing γ to be biased downward. After differencing, the error no longer contains u_i so there is no longer correlation between the regressors and the error term.[7]

Estimation of the difference Equation 6 takes advantage of an identifiable, exogenous source of variation in coverage – the FCIRA caused insurance coverage to increase more for some crops and regions than for other crops and regions. Insurance coverage grew more in some regions than others because the structure of the subsidy was such that it was more valuable in some regions as compared with others. In riskier regions, the subsidy was worth more because yields were more likely to drop below 50% of expected yields. Also, the subsidy was effectively

[6] In 1992, information on the number of days worked off-farm was not available in the ARMS survey, so we used the 1993 survey information. Values for the five intervals used were 20.2, 80.7, 128.3 and 184.7 for 1992 and 10, 72.4, 121.0 and 173.8 for 1997. This adjustment did not substantially change the significance or magnitude of the estimates.

[7] As an alternative to differencing, one can include fixed effects for each farm in Equation 1. Due to the large number of farms in the sample and the non-linear statistical methods that are employed, this approach was computationally infeasible.

greater for farms with higher average yields and those with higher valued crops (e.g., corn versus wheat).

The estimated effect of the insurance policy change $\tilde{\gamma}$ is unbiased provided that factors correlated with the change in insurance coverage did not simultaneously alter labour decisions. In fact, around the same time as FCIRA, another policy change occurred that might have caused changes in labour supply. The 1996 Federal Agricultural Improvement Reform Act (FAIR) dramatically altered the structure of agricultural income support payments. This Act, sometimes called the 'Freedom to Farm Bill,' decoupled most payments from farmers' current planting decisions. Prior to the FAIR Act, most government payments to farmers were tied to commodity prices, and farmers were required to limit current plantings to a share of historical plantings to qualify for payments. The FAIR Act lifted nearly all planting restrictions and decoupled payments from price levels. In effect, the Act scheduled lump-sum payments to land units based on pre-Act participation in government farm programmes. If the Act caused changes in off-farm labour decisions to change in a way that was correlated cross-sectionally with changes induced by FCIRA, it could bias the estimates.

To control for the effects of the FAIR Act, each farm's level of 1997 government farm payments is included as an explanatory variable in the vector X_{it}.[8] The level of these payments was determined in advance according to parameters laid out in the FAIR. The larger these payments, the more a farm was engaged in pre-1996 farm programmes, and the greater the effect of the policy change on income variability and insurance coverage, all else the same.

V. Data

Data on farm and operator characteristics are from the farm-level files of the 1992 and 1997 Census of Agriculture maintained by the National Agricultural Statistics Service (NASS) of the US Department of Agriculture. The census is conducted every five years and includes essentially all farms in the USA. Each farm operation receiving a census form is assigned a unique Census File Number (CFN). The CFN refers to the farm, ranch, or other agricultural activity controlled or operated by the person filing the census. To be sure that the same operator and operation are tracked over time the sample is limited to those operations in 1992 and 1997 having the same CFN and an having an operator whose age differed between 1997 and 1992 by five years.

Merging census records from 1992 and 1997 by CFN resulted in a panel data set with 2 083 386 observations. In order to exclude farms that were unlikely to be affected by the Federal Crop Insurance Reform Act, the sample was restricted to farms having a Standard Industrialization Code (SIC) classifying them as producers of a major insurable crop (wheat, corn, soybeans, 'cash grains' (oilseed and grain combination farms), cotton, or Irish potatoes), which resulted in 558 153 observations.

The sample was further restricted to operations for which land harvested in these insurable crops comprised at least 90% of the farm's total harvested land, resulting in 403 030 observations. Then a balanced panel was created by keeping only those operations observed in consecutive periods, resulting in 302 266 observations. The sample was further restricted to those operations with the same operator in 1992 and 1997 – as indicated by a change in the age of the operator by five years between censuses – resulting in 264 388 observations. Finally, observations with missing data were dropped to obtain a sample with 215 058 observations or 107 529 differences. As is discussed below: of the 107 529 farms with observations in both periods, 70 088 supplied no labour in both periods, or supplied 200 or more days of labour in both periods implying changes in their desired supply of off-farm labour is unobservable. Hence, the sample for which one can observe a (possibly censored) change in desired off-farm labour consists of 37 456 farms.

Table 2 reports summary statistics for the main variables used in the analysis for the sample. On average, farm operators worked 91.3 days in 1992, and 71.5 days in 1997. It is likely that this decline in off-farm labour supply reflects the ageing of the sample of operators between censuses. Operators were fairly evenly distributed over the 5-year age categories between 35 and 70 years, and had an average of 23.2 years of farming experience. About 98% of operators were male. Four value-of-production categories were constructed to allow for differential effects of coverage on labour supply. About 26% of farms produced less than \$25 000 worth of output, 38% produced between \$25 000 and \$100 000; 27% produced between \$100 000 and \$250 000 and 9% produced more than \$250 000 of output.

[8] Conservation Reserve Program (CRP) payments were not included in total government payments because enrolment in the programme is voluntary. In 1997, nearly all payments to farmers (net of CRP payments) were scheduled by the 1996 Act.

Table 2. Variable definitions and summary statistics for sample

Variable	Definition	Mean	Std. dev.
Lag days off	Number of days worked off-farm by operator, 1992	91.259	87.185
Δ days off	Change in number of days worked off-farm 1992–1997	−19.766	116.456
Lag insurance (1)	Insurance coverage (C^1) per acre harvested, 1992	2.632	3.790
Δ insurance (1)	Change in insurance coverage (C^1) per acre harvested, 1992–1997	2.649	4.934
Lag insurance (2)	Insurance coverage (C^2) per acre harvested, 1992	2.660	1.969
Δ insurance (2)	Change in insurance coverage (C^2) per acre harvested, 1992–1997	2.978	4.161
Lag insurance (3)	Insurance coverage (C^3) per acre harvested, 1992	2.660	1.969
Δ insurance (3)	Change in insurance coverage (C^3) per acre harvested, 1992–1997	2.799	2.188
Lag VOP <25	Value of production 1992 ($) < $25 000	0.257	0.462
Lag VOP 25–100	$25 000 < Value of production ($) 1992 ≤ $100 000	0.377	0.513
Lag VOP 100–250	$100 000 < Value of production ($) 1992 ≤ $250 000	0.275	0.472
Lag VOP >250	Value of production ($) 1992 > $250 000	0.092	0.305
Lag wage	Average county hourly wage per job 1992 (Bureau of Economic Analysis, December 2002)	6.707	1.281
Δ wage	Change in average county hourly wage per job 1992–1997	0.150	0.435
Gov_pay_acre 97	Total government payments per acre harvested in 1997, excluding Conservation Reserve Program payments	14.947	55.613
Experience	Years of farming experience of operator, 1997	23.220	12.795
Sex	Sex of operator: male = 1, female = 2	1.019	0.143
Age <35	Age, 1997 ≤ 35	0.071	0.272
Age 35–40	35 < age, 1997 ≤ 40	0.104	0.323
Age 40–45	40 < age, 1997 ≤ 45	0.138	0.365
Age 45–50	45 < age, 1997 ≤ 50	0.139	0.366
Age 50–55	50 < age, 1997 ≤ 55	0.123	0.348
Age 55–60	55 < age, 1997 ≤ 60	0.123	0.347
Age 60–65	60 < age, 1997 ≤ 65	0.121	0.346
Age 65–70	65 < age, 1997 ≤ 70	0.092	0.306
Age >70	70 < age, 1997	0.087	0.298
Lag wheat	SIC = 111, 1992	0.115	0.337
Lag corn	SIC = 115, 1992	0.413	0.521
Lag soybeans	SIC = 116, 1992	0.204	0.426
Lag cash grains	SIC = 119 (oilseed and grain combination farms), 1992	0.224	0.441
Lag cotton	SIC = 131, 1992	0.036	0.196
Lag potatoes	SIC = 134 (Irish potatoes), 1992	0.008	0.095
Observations		37 456	

Source: All variables from the Census of Agriculture, 1992 and 1997, unless indicated.

Measures of insurance coverage

A farm-level measure of insurance coverage is constructed that approximates insurance premiums per acre harvested. The approximation uses county-level information about crop insurance premiums obtained from the US Department of Agriculture Risk Management Agency (RMA). The reported premium equals farmer contributions plus government subsidies. The premium is based on farmers' expected indemnity and therefore provides a good measure of coverage.[9]

The study uses farm-level information on land use from the Census of Agriculture and county-average premiums per acre harvested from RMA to estimate individual operators' insurance coverage. Specifically, operator i's coverage per acre harvested in time t is defined as the weighted average premium per acre harvested for each crop in the county in which the farm is located, where the weights are given by the share of land each operator has in a particular crop:

$$C_{it}^{1} = \sum_{j} \left(\frac{P_{jt}^{c}}{A_{jt}^{c}} \right) s_{ijt} \tag{8}$$

where P_{jt}^{c} is the reported total RMA premium (farmer contribution plus government subsidy) and A_{jt}^{c} is the total land harvested for crop j in the county c in which farm i is located, and s_{ijt} is the share of land that farm i has in crop j at time t. Because coverage is expressed per acre *harvested* rather than per acre *insured*, the measure increases between censuses if,

[9] See the Risk Management Agency website for details about calculating the premium for specific crops and coverage levels: http://www.rma.usda.gov/

county-wide, farmers increase coverage on land already insured or enrol a greater share of cropland in the insurance programme at the same level of coverage.

There is an ancillary benefit to using county-level rather than farm-level measures of insurance coverage. At the farm level, any number of factors may cause a farm to simultaneously alter both insurance coverage and labour supply. Using county-level coverage measures effectively instruments the farm-level change in coverage with the county-level change in coverage. This removes idiosyncratic variance in coverage that is most likely to suffer from endogeneity problems.

Despite the use of county-level coverage levels for each crop, it is possible county-average premiums per acre harvested (P^c_{jt}/A^c_{jt}) are spuriously correlated with labour supply. This may occur if unobserved county-wide factors simultaneously affect both crop insurance and off-farm labour decisions, or if omitted county-wide factors are correlated with both decisions. To address this possibility, an alternative measure of insurance coverage is constructed that substitutes the average national premium per acre P^N_{jt}/A^N_{jt} for the county average in Equation 8:

$$C^2_{it} = \sum_j \left(\frac{P^N_{jt}}{A^N_{jt}}\right) s_{ijt} \tag{9}$$

Because total acres harvested in each crop changed relatively little between censuses at the national level (see Table 1) the change in the average national premium per harvested acre is due mainly to the exogenous policy change. Different growth rates in coverage still arise on different farms because farmers have different crop mixes and some crops were more subsidized than others due to the structure of the policy. This is evident in Table 1. Corn, for example, is a higher value crop than wheat, so holding yield variance constant, the subsidy embodied in catastrophic coverage is higher. Per-acre corn premiums thus grew by \$3.68 per acre compared with \$2.63 per acre for wheat.

Although aggregate changes in premiums per acre are plausibly exogenous, farm-level changes in crop shares s_{ijt} may remain endogenous. Some of these changes may have been caused by unobserved factors that jointly determine insurance and off-farm labour supply decisions. A third coverage measure similar to Equation 9 is therefore constructed except crop shares are held at their initial (1992) level:

$$C^3_{it} = \sum_j \left(\frac{P^N_{jt}}{A^N_{jt}}\right) s_{ij0}$$

Table 3. Total insurance coverage by value of production category and year

Value of production category	1992	1997
VOP < \$25 000	197.89	360.12
	(345.95)	(592.97)
\$25 000 < VOP ≤ \$100 000	846.53	1530.1
	(1097.0)	(1954.9)
\$100 000 < VOP ≤ \$250 000	1853.7	3524.9
	(2267.6)	(3797.8)
VOP > \$250 000	4461.8	7874.6
	(10 179)	(10 988)
Observations	37 456	37 456

Source: Census of Agriculture, 1992 and 1997 and Risk Management Agency. Standard deviations in parentheses.

With this measure, differences in coverage growth across farms stem from pre-FCIRA differences in crop mix across farms interacted with aggregate FCIRA-induced coverage growth. This measure is arguably more exogenous than the other two measures.

Table 2 reports the average insurance coverage per acre harvested in 1992 and the change in insurance coverage between 1992 and 1997. For the sample, the average coverage per acre harvested using definition 1 doubled from \$2.63 in 1992 to \$5.28 in 1997. The three measures of coverage had similar means in 1992, though measures 2 and 3 had less variation than measure 1. The average *change* in coverage was also similar for the three measures, though change in measure 3 had less variation than change in measure 2 or 3. The reduction in variance from measures 1 through 3 makes sense, as variance most likely to be endogenous has been removed. This variance reduction is analogous to that occurring with instrumental variables or two-stage least squares: the predicted value from the first stage has a smaller variance than the raw, un-instrumented variable.

Table 3 displays the total value of insurance coverage for farms in the sample in 1992 and 1997 by value of production category. The table illustrates that FCIRA resulted in a markedly larger absolute increase in total insurance coverage for large farms relative to small farms. The policy resulted in a \$162 increase in coverage for the smallest farm size category compared with an increase of \$684, \$1671, and \$3413 for the three larger farm size categories. Because larger farms had a greater absolute increase in insurance coverage, one expects operators of these farms responded more to the policy change than operators of smaller farms.

Table 4. Censored differences model of off-farm labour supply using alternative measures of crop insurance coverage

	(1)		(2)		(3)	
Parameter	Estimate	Std. err.	Estimate	Std. err.	Estimate	Std. err.
Intercept	149.15**	(67.19)	160.94**	(66.59)	248.13***	(67.60)
Δ insurance ∗ Lag VOP <25	−0.50	(0.78)	0.93	(0.86)	3.56*	(2.05)
Δ insurance ∗ Lag VOP 25–100	−0.86	(0.69)	−0.11	(0.75)	−0.49	(1.37)
Δ insurance ∗ Lag VOP 100–250	−3.36***	(0.73)	−2.12**	(0.89)	−10.04***	(1.46)
Δ insurance ∗ Lag VOP >250	−6.03***	(1.30)	−8.20***	(1.48)	−21.60***	(2.32)
Lag VOP <25	−190.32***	(8.62)	−202.95***	(9.22)	−246.33***	(11.56)
Lag VOP 25–100	−83.85***	(6.81)	−94.04***	(7.48)	−129.43***	(9.32)
Lag VOP 100–250	−32.40***	(6.69)	−43.39***	(7.50)	−57.95***	(9.23)
Experience	−0.55**	(0.18)	−0.54***	(0.18)	−0.50***	(0.18)
Sex = female	−68.10***	(12.49)	−67.25***	(12.50)	−66.72***	(12.47)
Δ wage	6.08*	(3.55)	6.22*	(3.56)	7.68**	(3.57)
Gov_pay_acre 97	0.11**	(0.05)	0.12**	(0.05)	0.11**	(0.05)
Age 35–40	−2.56	(7.06)	−2.67	(7.06)	−2.71	(7.05)
Age 40–45	−0.92	(6.84)	−1.27	(6.84)	−0.98	(6.83)
Age 45–50	15.91**	(6.98)	15.59**	(6.98)	15.35**	(6.96)
Age 50–55	13.28*	(7.37)	13.03*	(7.37)	12.70*	(7.35)
Age 55–60	−28.21***	(7.70)	−28.33***	(7.70)	−29.51***	(7.69)
Age 60–65	−74.40***	(8.21)	−74.70***	(8.21)	−75.27***	(8.20)
Age 65–70	−126.61***	(9.13)	−126.87***	(9.13)	−128.45***	(9.12)
Age >70	−109.96***	(9.85)	−110.24***	(9.86)	−113.67***	(9.86)
Lag wheat	−50.86**	(21.17)	−52.09**	(21.17)	−104.32***	(22.17)
Lag corn	−59.35***	(20.95)	−61.39***	(20.94)	−114.58***	(21.90)
Lag soybeans	−52.88**	(21.17)	−55.17***	(21.16)	−109.54***	(22.21)
Lag cash grains	−29.46	(20.92)	−31.72	(20.90)	−84.88***	(21.86)
Lag cotton	−28.15	(22.87)	−28.32	(22.83)	−52.08**	(22.83)
State dummies	Yes		Yes		Yes	
Log likelihood	−107 907		−107 909		−107 865	

Notes: Models (1)–(3) correspond to the three definitions of insurance coverage described in text. Number of observations = 37 456 (non-censored = 10 775, right censored = 11 288, left censored = 15 393).
*** significant at the 1% level; ** 5% level; * 10% level.

VI. Results

Table 4 presents model estimates in three columns corresponding to the three measures of insurance coverage defined in Section IV. Out of 107 529 farms, 70 073 had unobservable changes in their off-farm labour supply (because they supplied no labour in both periods, or supplied 200 or more days of labor in both periods) leaving 37 456 observations. Of these, 10 775 were uncensored, 11 288 were right-censored, and 15 393 were left-censored. Estimates were obtained using the QLIM procedure of SAS 9.1. To control for heteroscedasticity, the variance of the error in Equation 7 is modelled as a function of the parameters σ^2 and δ and z_i, a subset of the covariates: $E(\varepsilon_i^2) = \sigma_i^2 = \sigma^2(1 + \exp(z_i'\delta))$.

The three specifications produce similar parameter estimates and significance levels.[10] Operator's age in 1992 was significant in explaining the change in desired labour supplied off-farm between censuses. Relative to the farmers younger than 35 years old (the missing category), farmers between the ages of 45 and 55 increased their supply of labour over the five year period, whereas farmers over 55 years decreased their supply of labour. Farmers between 65 and 70 years reduced their labour supply more than did farmers in the oldest age category (over 70 years), probably because many in the 65–70 years bracket in 1992 entered retirement during the subsequent five-year period.

Using the third definition of coverage, an increase in the local wage rate between 1992 and 1997 of one dollar per hour increased off-farm labour supply by 12.3 days. Farmers worked an average of 81.4 days per year off farm in 1992 and 1997. Assuming they were paid at the average county wage of \$6.78 per hour, an *off-farm* labour supply elasticity of 0.64 is estimated. Compared with other estimates

[10] In addition to the variables included in the analysis, education would be expected to influence off-farm labour supply. Unfortunately, educational attainment is not recorded in the Agricultural Census.

of the labour supply elasticity, this estimate seems reasonable.[11]

The size of an operation was also correlated with the change in off-farm labour supply. Operators of the smallest farms saw the largest reduction in labour supplied between censuses, compared with operators of the largest farms (the missing category). In addition, the sex of the operator was important: compared with female operators, male operators decreased their supply of labour off farm more between 1992 and 1997. Farming experience was also associated with a reduction in the supply of labour – an additional year of experience reduced the change in supply of labour by 0.50 days. Government payments had a small, but statistically significant positive effect on labour supply. This result could reflect the 'decoupling' of payments from production associated with 1996 farm bill, which allowed farmers to continue to receive payments even if they switched from programme crops to less labour intensive non-programme crops.

State dummy variables were included in the regressions to control for differences in the rates of change in labour supply that may be correlated with unobservable regional factors, such as regional differences in changes in economic conditions. A joint test of these fixed effects indicates they are statistically significant.

Of central interest for this study, the change in insurance coverage was found to be associated with a statistically significant increase in the amount of labour supplied off farm for the largest two value-of-production categories for all three definitions of coverage. The labour response to a change in coverage for the largest value-of-production group is somewhat greater using coverage measure 3. An additional dollar in coverage per acre was estimated to reduce off-farm labour supply by 2.1–10.0 days for farms with value of production between \$100 000 and \$250 000 and by 6.0–21.6 days for farms with more than \$250 000 of production. These coefficients seem substantial given that a dollar-per-acre increase in coverage is equivalent to an increase in total premiums of \$697 and \$1677 for value-of-production categories 3 and 4, respectively, and implies exposure to risk significantly influenced farmers' labour decisions. The smallest farm size category showed a weakly significant positive response to the change in insurance coverage for the third measure of coverage. It is possible that the small-farm operators who experience the greatest increase in coverage were located in regions where the large-farm operators also had the greatest increase in coverage. These regions will see the largest declines in large-farm operator off-farm labour supply in response to the policy, which would presumably enhance the off-farm employment opportunities for the small-farm operators. The positive labour supply response of small operators, therefore, may reflect improved off-farm employment opportunities rather than a response to a change in their business risk.

Finally, Table 5 presents estimates of four alternative specifications of the censored regression model using the third definition of coverage. The table illustrates how the coefficients change as more controls are included in the regressions. The first column lists the coefficients from the third column of Table 4. The second column includes interaction effects between the lagged value-of-production categories and the age categories. The third column also includes interactions between the value-of-production and SIC categories, and between the age and SIC categories. The fourth column additionally allows for interaction between both the wage and government payments and the value-of-production, age and SIC categories. The table shows the estimated coefficients for the two larges size categories are robust to alternative model specifications. The coefficient for the smallest farm-size category, however, increases somewhat in magnitude and significance as additional controls are included.

VII. Conclusion

This study used an exogenous increase in Federal crop insurance subsidies induced by the Federal Crop Insurance Reform Act of 1994 as a natural experiment to identify the importance of risk in off-farm labour supply. Increases in crop insurance subsidies stemming from the Act induced greater participation in crop insurance programmes and expanded crop insurance coverage, which served to reduce farmers' financial risks. This study compared changes over time in the off-farm labour supply of farms that faced

[11] A survey of 50 labour market studies reported an average wage elasticity of −0.06 for men, and 0.94 for women (Hansson and Stuart, 1985). Women's labour supply is more elastic than men's in part because they work more in the home, allowing them to shift their allocation of labour from the home to the workplace without adjusting their hours of leisure. Similarly, farmers have a greater ability to adjust their supply of market labour than non-farmers do because farmers can shift their allocation of labour from the farm to the workplace without adjusting their leisure time. Hence, although almost all farm operators are men, it is not surprising that the estimated off-farm labour supply elasticity is closer to that of non-farming women than non-farming men.

Table 5. Censored differences model of off-farm labour supply under alternative model specifications

	(1)		(2)		(3)		(4)	
Parameter	Estimate	Std. err.	Estimate	Std. err.	Estimate	Std. err.	Estimate	Std. err.
Intercept	248.13***	(67.60)	317.96***	(74.74)	93.67	(101.10)	104.33	(100.10)
Δ insurance * lag VOP <25	3.56*	(2.05)	8.97***	(2.28)	8.61***	(2.31)	8.53***	(2.33)
Δ insurance * lag VOP 25–100	−0.49	(1.37)	2.19	(1.53)	2.09	(1.54)	2.04	(1.54)
Δ insurance * lag VOP 100–250	−10.04***	(1.46)	−12.70***	(1.69)	−12.31***	(1.67)	−12.14***	(1.68)
Δ insurance * lag VOP >250	−21.60***	(2.32)	−26.43***	(2.91)	−25.10***	(2.93)	−24.47***	(2.93)
Experience	−0.50***	(0.18)	−0.47**	(0.18)	−0.61***	(0.18)	−0.61***	(0.18)
Sex = female	−66.72***	(12.47)	−67.29***	(12.46)	−66.28***	(12.46)	−64.61***	(12.49)
Δ wage	7.68**	(3.57)	7.75**	(3.50)	8.07**	(3.48)	No	
Gov_pay_acre 97	0.11**	(0.05)	0.11**	(0.05)	0.11**	(0.05)	No	
State dummies	Yes		Yes		Yes		Yes	
Lag VOP dummies	Yes		No		No		No	
Age dummies	Yes		Yes		No		No	
Lag SIC (main crop) dummies	Yes		Yes		No		No	
Lag VOP * Age dummies	No		Yes		Yes		Yes	
Lag VOP * lag SIC dummies	No		No		Yes		Yes	
AGE * lag SIC dummies	No		No		Yes		Yes	
Δ wage * VOP interaction	No		No		No		Yes	
Δ wage * Age interaction	No		No		No		Yes	
Δ wage * lag SIC interaction	No		No		No		Yes	
Gov_pay_acre 97 * VOP interaction	No		No		No		Yes	
Gov_pay_acre 97 * Age interaction	No		No		No		Yes	
Gov_pay_acre 97 * lag SIC interaction	No		No		No		Yes	
Log likelihood	−107 865		−107 822		−107 660		−107 622	

Notes: Models (1)–(4) use the third measure of insurance coverage. Number of observations = 37 456 (non-censored = 10 775, right censored = 11 288, left censored = 15 393).
*** significant at the 1% level; ** 5% level; * 10% level.

different levels of insurance coverage before and after the policy change. This approach holds factors common to the farm household constant across time, and thereby controls for unobserved heterogeneity.

We find that increases in crop insurance significantly reduced the off-farm labour supply of operators of medium and large farms, but did not significantly alter the labour allocation decisions of operators of farms producing less than $100 000 of output. These results are found to be robust to changes in how crop insurance coverage was measured and to econometric model specification. The findings are generally consistent with earlier studies that found risk – measured by historical variation in farm income – inversely correlated with off-farm labour supply. The results elaborate on those studies by showing that labour responses to risk-reducing policy changes are not consistent across farm-scale classes. It is found that larger farms – those with the most insured acres – had the greatest off-farm labour response to the large change in the cost of bearing agricultural risk caused by insurance policy change. For smaller farmers, for whom agriculture represents a relatively small portion of total income, the policy change either did not change or slightly increased off-farm labour supply.

The results suggest that despite large government expenditures on crop insurance and other risk mitigating agricultural programmes, farmers alter their production decisions, in particular their allocation of labour on and off-farm, to cope with income risk. By reducing farmers' dependence on such potentially costly risk-coping mechanisms, the findings suggest that crop insurance programmes may improve economic efficiency. However, the results also highlight possible unintended consequences of crop insurance policies for rural labour markets. Though not explored here, similarities between farm and non-farm small businesses suggest non-farm sole proprietors may also respond to business risk by altering their labour market decisions.

Acknowledgements

The authors are economists at the Economic Research Service, US Department of Agriculture. The views expressed are those of the authors and do not necessarily correspond to the views or policies of the Economic Research Service or the US Department of Agriculture.

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